HOW MUCH FAVORABLE SELECTION IS LEFT IN MEDICARE ADVANTAGE?

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ABSTRACT
The health economics literature contains two models of selection, one with endogenous plan characteristics to attract good risks and one with fixed plan characteristics; neither model contains a regulator. Medicare Advantage, a principal example of selection in the literature, is, however, subject to anti-selection regulations. Because selection causes economic inefficiency and because the historically favorable selection into Medicare Advantage plans increased government cost, the effectiveness of the anti-selection regulations is an important policy question, especially since the Medicare Advantage program has grown to comprise 30 percent of Medicare beneficiaries. Moreover, similar anti-selection regulations are being used in health insurance exchanges for those under 65.

Contrary to earlier work, we show that the strengthened anti-selection regulations that Medicare introduced starting in 2004 markedly reduced government overpayment attributable to favorable selection in Medicare Advantage. At least some of the remaining selection is plausibly related to fixed plan characteristics of Traditional Medicare versus Medicare Advantage rather than changed selection strategies by Medicare Advantage plans.

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I. Introduction

The health economics literature contains two types of selection models. One focuses on the incentives of risk-bearing insurers to attract good risks by tailoring the insurance options they offer to appeal to good risks (Rothschild and Stiglitz 1976; Newhouse 1996; Hendren 2013). Inefficiency in these models takes the form of insurance contracts with the wrong mix of premiums and benefits and/or of resources devoted to particular benefit areas (Frank, Glazer, and McGuire 2000; McGuire et al. 2014). As is well known, an

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equilibrium may or may not exist in this type of model depending on assumptions about the distribution of risks in the population and insurers' behavior (Wilson 1977; Dubey and Geanakoplos 2002; Breyer, Bundorf, and Pauly 2012). Risk adjustment can improve the efficiency of contracts, but some inefficiency remains if risk adjustment is set at the expected cost of an enrollee with given observable characteristics, as is typically the case (Glazer and McGuire 2000).

A second selection model, which is nested within the first, holds insurance contracts fixed, focuses on consumers, and makes the assumption that the demand for contracts with more generous reimbursement is positively correlated with anticipated health-care use. If premiums are the same for every person with given observable characteristics who chooses a given contract, as is generally the case, and if consumers have a choice of contracts, the resulting selection implies an equilibrium with too few people choosing more generous contracts (Feldman and Dowd 1982; Cutler and Reber 1998; Einav, Finkelstein, and Cullen 2010; Einav and Finkelstein 2011).[1] In the extreme, a death spiral can ensue, and more generous plans can disappear altogether (Shore and Bertko 1999; Yegian et al. 2000; Weinberg and Kramer 2011).

This paper studies trends in selection in Part C of Medicare, or Medicare Advantage (MA), in light of these theoretical perspectives. The amount of selection in MA is an important policy question, since 30 percent of Medicare beneficiaries now enroll in it (Gold et al. 2014), and historically the health economics literature has convincingly demonstrated that in the 1990s those newly enrolling in MA were relatively good risks and those newly disenrolling may have been relatively poor risks (Congressional Budget Office 1997; Morgan et al. 1997; Cutler and Zeckhauser 2000; Glied 2000; Newhouse 2002; McGuire, Newhouse, and Sinaiko 2011; Breyer, Bundorf, and Pauly 2012). As a result of this favorable selection, the Congressional Budget Office (CBO) estimated in 1997 that Medicare paid 8 percent more for beneficiaries enrolled in MA than it would have paid if the same enrollees had enrolled in Traditional Medicare (TM), that is, if MA had not existed, even though at that time Medicare only paid 95 percent of the average adjusted TM payment in the beneficiary's county of residence for MA enrollees.

As explained below, in the mid-2000s Medicare adopted policies intended to reduce overpayment attributable to favorable selection in MA. Surprisingly, Brown et al. (2014), interpreting their findings within the first model of selection mentioned above, concluded that insurers responded to the policy changes by changing their selection mechanisms such that the degree of overpayment in MA actually increased. In other words, Brown et al. concluded that the policy changes were counterproductive.

This paper reexamines the question of how selection has changed in response to Medicare policy. In contrast to Brown et al., we reach the conclusion that overpayment attributable to favorable selection in MA has fallen markedly and may now be at a level that falls below a concern for Medicare payment policy. The effectiveness of the new anti-selection instruments is an important issue not only for the Medicare program

1 Sufficient risk aversion among good risks can lead to favorable selection in these models (Finkelstein and McGarry 2006).

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but also for the Affordable Care Act's health insurance exchanges, because most of the tools used in the Medicare program to address selection are also being used in the exchanges.

II. Background on the Medicare Advantage Program

Until 2006 insurers participating in the MA program agreed to accept a fixed per enrollee payment that was set annually. This take-it-or-leave-it payment was a function of spending by the average TM beneficiary in the MA beneficiary's county, although the nature of the function changed over time (McGuire, Newhouse, and Sinaiko 2011). In return, MA plans were to provide benefits at least as comprehensive as the "public option" of TM (McGuire, Newhouse, and Sinaiko 2011).[2] Starting in 2006 and continuing to the present, each insurer submits a bid rather than simply agreeing to a payment set by Medicare. But a modified version of the pre-2006 administered price system lives on; instead of the insurer's receiving an administered price, the beneficiary receives an analog to a voucher that equals the average amount Medicare would have paid to the insurer if the pre-2006 take-it-or-leave-it regime had continued. The analog is called the benchmark, and if the insurer's bid is less than the benchmark, as almost all bids are, the beneficiary receives 50–70 percent of the difference between the bid and the benchmark as additional services, lower cost sharing, or lower premiums.[3] The government keeps the other 30–50 percent. In the few cases in which the bid exceeds the voucher, the beneficiary pays the entire increment. Importantly, TM is not part of this bidding system. For more detail see Newhouse and McGuire (2014).

MA insurers compete not only with each other but also with TM through the terms of the contracts or insurance plans that they offer beneficiaries. At a high level of generality the attraction of MA relative to TM for a beneficiary without supplementary insurance (Medigap) is much less cost sharing at the point of service and often additional covered services such as vision or dental.[4] For a beneficiary with a Medigap plan the attraction

2 The requirement that benefits be at least as comprehensive as TM meant (and means) that MA plans had to cover any service covered by TM, for example, physician visits. Further, the actuarial value of any beneficiary out-of-pocket payments in the MA plan, both premiums and cost sharing, could not (and cannot) exceed the actuarial value of the cost sharing in TM. MA participants were (and are) required to be enrolled in Part B and pay the Part B premium.

3 The percentage the beneficiary receives varies with the rated quality of the plan; the higher the quality, the higher the percentage.

4 A beneficiary enrolling in TM has the option of purchasing an individual Medigap plan to cover most or all of the cost sharing in TM, in effect converting the additional cost sharing in TM relative to MA to an additional premium relative to MA. Around 30 percent of beneficiaries have a supplementary plan that is partly or fully subsidized by a prior employer, although the subsidy may be structured as a lump sum subsidy with MA plans as a choice. Around 18 percent of all Medicare beneficiaries (and 13 percent of elderly beneficiaries) are eligible for Medicaid, which historically functioned like a supplementary plan to TM, although many states are currently enrolling Medicaid beneficiaries in MA with the ability to opt out to TM (Medicare Payment Advisory Commission 2013).

is in not having to pay the Medigap premium and, very likely, lower cost sharing as well, especially for drugs (Medicare Payment Advisory Commission 2014). In return, the beneficiary agrees to change the terms at which he or she can use certain physicians or hospitals; if a given physician or hospital is not in the MA plan's network, the beneficiary's out-of-pocket cost to use that physician or hospital would likely be considerably above what the beneficiary would pay in TM.

To finance the lower cost sharing they typically offer, MA plans must reduce the quantity of services they pay for, either through medical management of chronic diseases (e.g., attempting to increase compliance with prescribed medication and thereby lowering the likelihood of hospitalization), choosing physicians to be in their network who practice conservatively, or with the mechanism of interest to us, offering networks of providers and drug formularies[5] and configuring cost sharing in ways that will attract good risks, net of any risk-adjustment payments.[6] Risk-adjustment payments transfer monies from plans whose enrollees have lower expected cost to plans with enrollees with higher expected cost. In theory, risk adjustment lessens plans' incentives to attract the healthy and deter the sick and thus decreases distortions in service offerings (Glazer and McGuire 2000). To the degree risk adjustment is imperfect, however, MA plans' incentives resemble those of insurers in the first type of selection model referred to above.

Medicare requires that plans accept all beneficiaries who want to join, so to obtain a favorable risk mix, plans must influence who wants to join them in the first place, as in the first type of selection model. Their main instruments for doing so are their choice of networks of providers, covered benefits, and structure of drug formularies, as well as their marketing and choice of geographic area in which to operate (Frank, Glazer, and McGuire 2000; Cao and McGuire 2003).

The second type of selection model, although usually posed in the context of an employer offering plan alternatives to its employees, is also relevant to selection between TM and MA. Not surprisingly, those with more illnesses use more providers on average than those with fewer, as we show below. These persons may also have more complex diseases in unobserved ways and as a result could prefer TM's freedom of choice of provider (almost all providers participate in TM), since the more providers the beneficiary uses, the greater the chance that one or more of them is not in an MA plan's network. Networks, of course, are inherent in MA, and freedom of provider choice is inherent in TM. If selection of this type is going on, beneficiaries in MA plans will tend to be "healthier" and less costly

5 Drug formularies (analogous to provider networks) specify the drugs covered and copayments required, typically containing four "tiers" of coverage, with generics on the first tier and lowest copayments, followed by a tier with "preferred brands," a third tier with "non-preferred brands," and a specialty tier with coinsurance for very expensive drugs. Medicare regulates drug formularies offered but plans have some discretion about which drugs to include and the degree to which they are covered. MA beneficiaries obtain Part D coverage from their plan, although coverage is voluntary.

6 In principle, MA plans could also finance the lower cost sharing by paying lower unit prices than TM. These prices are not public information, but we are told by plan actuaries that the unit prices MA plans pay generally equal or slightly exceed TM unit prices.

than those in TM, but better regulation cannot address it given the current structure of the Medicare program.[7]

Medicare employs several instruments to mitigate plans' selection incentives.[8] First, as already alluded to, it risk adjusts the amounts it pays MA plans by making budget neutral transfers among MA plans according to certain observed characteristics of each plan's mix of enrollees. TM, however, is not part of this transfer system. Up until 2000, the algorithm to determine a beneficiary's expected cost only accounted for demographic variables such as age, gender, Medicaid eligibility, whether the beneficiary was institutionalized, and the beneficiary's county of residence. These variables were used to predict spending by TM beneficiaries for individuals in MA plans with given characteristics. For example, a plan that hypothetically enrolled only 65 to 69-year-old males in a given county would have been paid less than a plan that hypothetically enrolled only 70 to 74-year-old males in the same county. Second, Medicare regulates network and formulary adequacy; a plan's network, for example, must include a sufficient number of oncologists. Third, Medicare regulates a plan's choice of geographic area in which to operate; generally a plan must either operate in an entire county or not enter the market in that county; it requires special approval to choose particular zip codes within a county in which to operate. Fourth, as mentioned above, Medicare constrains the amount of cost sharing; the actuarial value of a plan, or the percentage of an average beneficiary's cost that the contract covers, must equal or exceed that of TM. Thus, a plan's ability to attract good risks by offering even higher cost sharing and lower premiums than TM, a standard kind of selection tool in unregulated insurance markets, is limited. In practice this actuarial value constraint is rarely binding, because, as noted above, MA plans typically offer much lower cost sharing than TM. Finally, Medicare has an overarching authority to refuse to contract with a plan that it deems is engaged in selection.

As a result of the favorable selection into MA in the 1990s described in the introduction, Medicare took two additional steps in the past decade to decrease it. First, it incorporated diagnostic information into its risk-adjustment scheme. Medicare's risk-adjustment method effectively sets a relative supply price for each plan enrollee as a function of the relative spending to treat an enrollee with similar observable characteristics in TM. Starting in 2000 Medicare began to adjust its then fixed payment to plans using information on inpatient diagnoses, but it gave that system only 10 percent weight; the other 90 percent of the weight continued to be on the old demographic risk-adjustment system.[9] This mixed, but mostly demographic, system continued until 2004, when Medicare began a

7 Newhouse et al. (2013) documented that the current risk-adjustment system undercorrects for risk among those with multiple comorbidities. In principle, this could be addressed by including more disease-disease interactions in the current risk-adjustment system, but sample sizes undoubtedly limit the number of feasible interactions. Thus, in practice feasible risk-adjustment systems are likely to offer incentives to select against those with infrequent combinations of multiple comorbidities.

8 Employers play this role in the commercial market by choice of plans to offer employees and by their subsidy policy towards employee share of premiums, but employers rarely use a formal risk-adjustment system such as that used by Medicare (Keenan et al. 2001).

9 The small weight on the inpatient diagnosis-based system was to mitigate an incentive to hospitalize a beneficiary just to record a diagnosis.

transition to a system that incorporated diagnostic information from both inpatient and outpatient settings (Pope et al. 2004). That system, which Medicare continues to use, is called CMS-HCCs or just HCCs, and in 2004 it had 30 percent weight in determining payment, in 2005 50 percent weight, in 2006 75 percent weight, and the transition was complete in 2007. From 2004 to 2006 the remaining weight was on the old system. The HCC risk-adjustment formula uses indicators for diagnoses recorded on claims in the previous year along with a limited number of interactions to predict costs in the current year.[10]

The new system was a major change from the earlier demographic adjustment system; it raised the explained variation in annual individual cost from around 1 percent to around 11 to 12 percent (Pope et al. 2004). Thus, under the HCC system an MA plan would be reimbursed considerably more for a 75-year-old woman living in the community with congestive heart failure (HCC80) than for a 75-year-old woman living in the same community with breast cancer (HCC10), whereas under the earlier demographic system it would have been reimbursed the same amount (Pope et al. 2004). Like the old system, however, the HCC risk-adjustment system simply makes budget neutral transfers among MA plans.

The second step Medicare took to reduce selection was to make it more difficult for beneficiaries to move between MA and TM. Prior to 2006 beneficiaries could move from MA to TM at the end of any month, in contrast to commercial insurance where an individual or an employer typically chooses a health insurance contract for an entire year during an annual open enrollment period. The ability of Medicare beneficiaries to change from MA to TM monthly was initially seen as a beneficiary protection against underservice by or dissatisfaction with an MA plan but it facilitated selection, since a beneficiary in MA with a new diagnosis who wanted to use an out-of-network physician to treat the problem could move to TM almost immediately.

Starting in 2006, however, a beneficiary who chose MA was locked in to the chosen plan for the last six months of the calendar year, but could still change plans monthly during the first half of the year. Beginning in 2007 the lock-in period was the last nine months and in 2011 it was extended to the last 10.5 months of the year. Nonetheless, by allowing a plan change in the first six weeks of the year Medicare continues to remain less restrictive than commercial insurance. Also the lock-in does not apply to the minority of Medicare beneficiaries also eligible for Medicaid, the dual eligibles.

Coincident with these changes, Medicare introduced a new type of MA plan, a Special Needs Plan (SNP). SNP plans were limited to certain beneficiaries. Three different kinds of SNPs were introduced, one for dual eligibles (those also eligible for Medicaid), one for the institutionalized, and one for those with certain chronic diseases.[11] Finally, another type of MA plan, Private Fee-for-Service (PFFS), was authorized in 1997 legislation and entered the market in 1998. PFFS plans were open to all beneficiaries, but attracted few enrollees

10 The HCC model also uses as adjusters age-gender cells, whether the beneficiary is eligible for Medicaid, and whether the original reason for Medicare eligibility was disability.

11 Of the various types of SNPs, the one for the dual eligibles, dominated numerically; 83 percent of SNP enrollees were in those plans (Medicare Payment Advisory Commission 2007).

before 2004, after which time they grew rapidly (McGuire, Newhouse, and Sinaiko 2011). We discuss PFFS plans further below.

III. Prior Estimates of Selection Effects

Virtually all the historical evidence on selection in Medicare and much of the more recent evidence compares use of medical services among those currently in TM who subsequently switch to MA with those who remain in TM. The limitation to those in TM reflects the availability of TM data and the lack of analogous MA data. In addition to examining those in TM who switch to MA, the literature also often compares those in TM who recently left MA with those who were always in TM.[12] These comparisons are typically adjusted for age, gender, Medicaid status, employment status, and county, the variables which were the principal risk adjusters for plan payment in the 1990s.

The early literature, using data from the 1990s, found that those enrolling in MA were relatively low utilizers while they were enrolled in TM immediately prior to their enrollment in MA, when compared with those who remained in TM and did not switch to MA. Data on the use of the much less numerous group disenrolling from MA were mildly conflicting; some data showed the disenrollees to be relatively high users after moving to TM when compared with the group that had remained in TM, and other data showed the use of the two groups was similar. In short, these data suggested that good risks were enrolling in MA and bad risks might be disenrolling (Physician Payment Review Commission 1996; Medicare Payment Advisory Commission 2000). It was on the basis of such comparisons between switchers and stayers that the CBO determined that the MA program added to Medicare program cost.

There are three problems with inferring the amount of selection from such studies. The first two potentially arise from status quo bias or switching costs, the tendency of beneficiaries to remain in the same plan year after year (Samuelson and Zeckhauser 1988). First, only a small percentage of beneficiaries switch from TM to MA each year, as we show below. As a result, even if those who switch exhibit the expected type of selection, the characteristics of the much larger stock of beneficiaries who remain in their plan types year after year could differ substantially from the characteristics of those who switch.

Second, and related to the first point, even if those few who switch differ in their unobserved characteristics affecting health or health-care use at the time of the switch, initial differences appear to regress to the mean. For example, in 1998 the age-sex-Medicaid adjusted mortality rate among those who were enrolled in MA for less than a year was 21 percent less than those who remained in TM, but for those who had been in MA five years or more, the difference was only 11 percent (Newhouse et al. 2012). We note in passing that these mortality data support the inference that in the 1990s there was favorable selection into MA since it is implausible that any causal effect of MA on mortality could be

12 Selection, of course, cannot be identified by a simple contemporaneous comparison of use in TM and MA, since differences could arise from selection or differences in medical management, the intensity of which is unobserved, as well as differences in cost sharing or in covered services, which may be difficult to control for.

21 percent or even 11 percent. These mortality differences, of course, apply to the stock of beneficiaries in each type of plan, not just the small flow between them.

Third, one would like to know the magnitude of selection net of those who switch into and out of MA, but inconsistent data interfere with such a calculation.[13] In the period we are examining, however, many fewer persons, only about a fifth as many, switch out of MA into TM as switch into MA from TM (Newhouse et al. 2013); hence we focus here on those switching into rather than out of MA.[14]

More recent data, however, show less evidence of favorable selection than the data from the 1990s did. By 2008 the difference in age-sex-Medicaid adjusted mortality among those enrolled in MA for less than a year had fallen from its 1998 value of 21 percent to 13 percent and the difference among those enrolled five years or more in MA, instead of 11 percent, was an insignificant 1 percent less, again consistent with regression toward the mean.[15]

Other data in addition to the mortality data imply selection behavior changed in the past decade. Using the HCC algorithm one can compute a risk score for each beneficiary, which is proportional to the average spending in TM conditional on observable characteristics including diagnosis.[16] The average risk score changed substantially among those switching into MA over the 2004–08 period relative to those remaining in TM. In 2003 those in TM switching into MA in 2004 had risk scores that were about 10 percent less than those who remained in TM, consistent with the earlier studies of favorable selection into MA; by 2008 the 10 percent figure had fallen to 3 percent (Newhouse et al. 2012).[17] This period coincided with the introduction of the CMS-HCCs and the lock-in.

These risk score data, of course, are subject to the caveat noted above that the characteristics of the switchers may not closely represent those of the stock of beneficiaries. McWilliams, Hsu, and Newhouse (2012), however, compared self-rated health status and

13 The difficulty in reaching a summary value that accounts for those switching in both directions is that the risk score for those switching into MA must be based on lagged diagnoses, since historically diagnostic data have not been available for MA enrollees. For the same reason among those disenrolling from MA the risk score must be based on concurrent diagnoses because lagged diagnoses are not available; as a result, these risk scores are not comparable. Nonetheless, the difference in numbers between those switching into MA relative to those switching out of MA means the characteristics of those switching in, the group we focus on, would likely dominate any summary value of selection.

14 This decision means our conclusions on selection may be somewhat optimistic because those disenrolling from MA into TM in the more recent period do appear sicker than those who remain in TM continuously (Newhouse et al. 2012).

15 The small difference of 1 percent could potentially be due to better medical management in MA plans.

16 The average risk score of an MA plan's enrollees is used to determine the plan's risk-adjustment payment.

17 These results are consistent with the findings of Brown et al. (2014) in their Figure 2. Brown et al. assert that the upper part of the risk score distribution was unprofitable; however, this assertion is based on TM spending patterns that do not replicate in MA and hence does not necessarily hold for MA enrollees. Newhouse et al. (2013) makes use of cost data by HCC from two MA plans; although in one plan the profitability of a beneficiary fell monotonically with the risk score, consistent with Brown et al., that was not the case in the other plan. Brown et al. interpret their results as consistent with the first model of plan selection; however, we interpret them as consistent with the second model of plan selection, as we describe below.

utilization among the stock of those in MA and those in TM over the 2001–07 period. Like the mortality data between 1998 and 2008, these comparisons showed a striking change. Whereas the proportion of beneficiaries who rated their health as fair or poor, the two worst categories, was over 20 percent less in MA than in TM in 2001–03, by 2006–07 it was only 5 percent less and one could not reject the null of no significant difference.[18] Consistent with the self-rated health changes, total utilization in MA was about 18 percent less in 2001–03, but only about 8 percent less in 2006–07. Drug fills, which in 2001–03 were about 5 percent less in MA than TM, were more than 10 percent greater by 2006–07. Comparisons of changes in the utilization of medical services over this period could reflect changes in medical management techniques as well as changes in selection, but any improvement in medical management techniques, the likely direction of any change, would mean the inference of reduced selection was understated.

IV. But What Happened to the Profitability of Plans?

Although the earlier literature suggested that in the 1990s plans benefited financially from favorable selection, the results just cited do not answer the question of whether plans changed their selection mechanism as the new HCC risk adjustment was phased in such a way that they maintained or enhanced their profit. After the implementation of the HCC system, plans' profit-maximizing strategy changed to select low utilizers within an HCC category rather than simply select low utilizers within an age-sex category (Brown et al. 2014). Brown et al.'s empirical evidence leads them to conclude that plans did change their behavior in exactly this fashion. In the remainder of this section we describe Brown et al.'s basis for this claim. We then reestimate their key equation on a much larger data set and a more targeted population.

Brown et al. (2014) estimated the following two equations using data on choices made by those in TM from 1994 to 2006 about whether to switch to MA in the following year. The two equations compare the risk score and Medicare spending, respectively, for persons in TM who switch from TM to MA for some part of the following year with those who do not switch to MA and remain in TM. The question of interest is whether those TM enrollees who opt for MA at the end of 2003 or in subsequent years, and so start MA enrollment between 2004 and 2007, differ in their use (relative to stayers in TM) from those opting for MA in the 1994–2002 period. The sample Brown et al. use comes from Medicare claims that are linked to the Medicare Current Beneficiary Survey (MCBS) and is limited to persons who were in TM in year t since no analogous data are available for those enrolled in MA.

$$\begin{aligned} \mathit{Risk\ Score}_{it} = {} & \alpha_0 + \alpha_1 \mathit{Fraction\ of\ Next\ Year\ Spent\ in\ MA}_{i,t+1} \\ & + \alpha_2 \mathit{Fraction\ of\ Next\ Year\ Spent\ in\ MA}_{i,t+1} \\ & \times \mathit{After\ 2002}_t + \alpha_3 \mathit{Year\ Dummies} + \epsilon_{it} \qquad (1). \end{aligned}$$

18 Even in 2006–07, however, about 10 percent more MA beneficiaries said their health had worsened relative to the prior year.

$$\begin{aligned} Gov\ \$_{it} = \beta_0 &+ \beta_1 Fraction\ of\ Next\ Year\ Spent\ in\ MA_{i,t+1} \\ &+ \beta_2 Fraction\ of\ Next\ Year\ Spent\ in\ MA_{i,t+1} \\ &\times After\ 2002_t + \beta_3 Year\ Dummies + \beta_4 Risk\ Score_{it} + \varepsilon_{it} \qquad (2). \end{aligned}$$

In both equations *Risk Score*$_{it}$ is the CMS-HCC value for beneficiary i in year t, and in the second equation *Gov* $\$_{it}$ is TM spending on beneficiary i in year t, including any out-of-pocket spending by the beneficiary. *After 2002* takes the value one for the years 2003 and later and is zero otherwise.

It is easiest to interpret these equations if one simply assumes that those who switched to MA spent the entire next year in MA so that the fraction of the next year spent in MA is one for those who switched and zero for those who did not. Then for those who switched into MA for 1995, the predicted risk score is $\alpha_0 + \alpha_1$, whereas for those who remained in TM it is just α_0. For subsequent years one simply adds the coefficient of the year dummy to both groups, with an additional effect for the MA group in the *After 2002* years. Consistent with the results from the literature cited above showing favorable selection in the 1990s, Brown et al. (2014) find α_1 negative and highly significant in eqation 1, indicating that those who switched to MA had a lower risk score when in TM and presumptively spent less. Newhouse et al. (2012) found a similar result.

The results of estimating equation 2 are of greater interest since it not only predicts total Medicare spending but also controls for the risk score, which began to be used to adjust payments to plans starting in 2004 as described above. *Gov* $\$_{it}$, the total spending by TM beneficiaries in 1994 who switched to MA in 1995 is $\beta_0 + \beta_1$, and for those who switched starting in 2003 it is $\beta_0 + \beta_1 + \beta_2$ plus a year fixed effect that applies to both switchers and stayers and differences out in comparing the two periods. Brown et al.'s (2014) test of whether the spending for those who chose to switch into MA in the years 2003 and later (and so were enrolled in MA in 2004 and later) differed from those who chose to switch in prior years is whether one can reject the hypothesis that $\beta_2 = 0$, since the transition to the HCC system began in 2004. They estimate β_2 to be negative with a t-statistic of 2.02. Since risk score is controlled for, Brown et al. interpret this as indicating that plans changed their selection mechanisms such that those within an HCC who were spending less in TM the year before they joined an MA plan were differentially attracted to MA.

Brown et al.'s (2014) use of claim files linked to the MCBS severely limits their sample size because the MCBS has only about 16,000 total observations per year. The share of the 16,000 observations in TM varies over the time period, but averages around 90 percent in the pre-HCC period and somewhat less in the post-period. Of that 90 percent about 1–4 percent switch into MA in a given year, or only about 150 to 600 people per year, which is likely why Brown et al. pooled the nine years before the introduction of the HCCs to compare with the four subsequent years. As we described earlier, the incentives for plans to select within HCC varied by year during the 2004–07 period because these four years encompassed the transition to the HCC method. To obtain greater precision we reestimated Brown et al.'s equations using the much larger 20 percent random sample of Medicare claims from each calendar year in the 2001–11 period.

TABLE 1. Beneficiaries by age and original reason for eligibility

	Total	Original reason for eligibility = aged		Original reason for eligibility = disability, end-stage renal disease, or Alzheimer's			
		Age 65+		Age 65+		Age <65	
2001	5,760,821	4,563,388	79.2%	422,207	7.3%	775,226	13.5%
2002	5,959,061	4,701,352	78.9%	442,511	7.4%	815,198	13.7%
2003	6,095,807	4,779,197	78.4%	458,698	7.5%	857,912	14.1%
2004	6,143,162	4,781,953	77.8%	465,479	7.6%	895,730	14.6%
2005	6,129,633	4,740,059	77.3%	467,811	7.6%	921,763	15.0%
2006	5,974,412	4,607,287	77.1%	459,927	7.7%	907,198	15.2%
2007	5,887,761	4,506,823	76.5%	460,829	7.8%	920,109	15.6%
2008	5,820,956	4,427,286	76.1%	465,482	8.0%	928,188	15.9%
2009	5,871,707	4,429,056	75.4%	476,663	8.1%	965,988	16.5%
2010	5,958,291	4,462,672	74.9%	498,891	8.4%	996,728	16.7%

Notes: The following types of beneficiaries were excluded from the sample: newly eligible beneficiaries since no prior claims information was available with which to compute a risk score and beneficiaries who did not have 12 months of continuous enrollment in TM, both Parts A and B, in the prior year. In addition, beneficiaries who switched into cost MA plans or Special Needs Plans were excluded.

Brown et al. (2014) include all TM enrollees in their sample, but we limit our sample to those whose original reason for Medicare eligibility was old age, thus excluding beneficiaries who became eligible before age 65 for reasons of disability or who have end-stage renal disease or Alzheimer's. Table 1 shows that the group whose original reason for eligibility was old age comprises 75–79 percent of Medicare beneficiaries. To not muddy comparisons across years by differences in populations, we also exclude the dual eligibles and the institutionalized. Dual eligibles refer to Medicare beneficiaries who are also currently eligible for Medicaid. As explained above, the 2003 Medicare Modernization Act created separate Medicare Advantage Special Needs Plans (SNPs) for the duals, the institutionalized, and those with certain chronic conditions as part of the MA program.[19] Thus, legislative authorization for these plans did not exist until 2003, so the menu of choices for the duals and the institutionalized expanded coincident with the introduction of the HCCs. Moreover, the lock-in provisions that took effect in 2006 did not apply to

19 The introduction of such plans could have changed selection patterns among the dual eligibles and institutionalized. These plans first entered the market in 2004, but there were only 11 such plans in that year, a number that grew to 125 plans in 2005. The first enrollment data we have found are for 2006, when there were 276 SNP plans with 541,000 enrollees out of a total of 6.9 million MA enrollees. In 2006, 226 of the 276 SNPs were for dual eligibles, and they had 440,000 enrollees (6 percent of all MA enrollees); there were also 37 plans for the institutionalized with 20,000 enrollees (Medicare Payment Advisory Commission 2007).

duals, and the institutionalized are paid with a separately estimated risk-adjustment formula. Finally, nearly half of those who were originally eligible for reasons of disability are also dual eligible and thus would have been excluded even if we had not limited the sample to those eligible for Medicare for reasons of age (Medicare Payment Advisory Commission 2013).

One type of plan that flourished between 2006 and 2010 was the Private Fee-for-Service (PFFS) plan. Although this plan was classified as and paid like an MA plan, its characteristics are similar to TM. In particular, PFFS plans in that period did not have networks, so that enrollees could see any provider who saw Medicare patients without a differential out-of-pocket payment; the provider was paid at TM rates.[20] Like TM but unlike other MA plans, there was no medical management.[21] PFFS plans only had 80,000 enrollees in 2005 (and less than half that in prior years), but that figure grew to 800,000 in 2006, 1.5 million in 2007, 2.3 million in 2008, 2.4 million in 2009, and then began to decline because of a requirement that PFFS plans have networks that took effect in 2011.[22] In this period PFFS dominated TM for many beneficiaries (McWilliams et al. 2011). Because PFFS plans both differ from HMO and PPO plans and also because they attracted many switchers in some years in our sample but not others, we will show results including and excluding those who switched to PFFS plans.

Like Brown et al. (2014), our sample includes only those who were enrolled in both Parts A and B of Medicare in the prior year and who were enrolled in Medicare on January 1 of the succeeding year. This latter restriction is necessary to allow us to identify those who switched from TM to MA, but it excludes those who died in the prior year. It also excludes the relatively small numbers of beneficiaries in cost-based MA plans, because the reimbursement for such plans is based on their medical costs and is not fixed as it is for most MA plans.[23] Cost-based plans enrolled 10 percent of enrollees in 1996, but only 4 percent in 2014 (Prospective Payment Assessment Commission 1997; Centers for Medicare and Medicaid Services 2014). Finally and importantly, risk scores for new Medicare enrollees are based only on demographic variables because newly eligible Medicare beneficiaries have no prior claims information. Lacking diagnostic information for this group with which one could compute an HCC score, we also exclude them.[24]

Before proceeding to our results, we summarize our improvements on Brown et al. (2014). First, we use a much larger sample of beneficiaries. Because of a relatively small sample, Brown et al. had to pool the years from 1994 to 2002 and compare selection in those years with selection during the pooled years from 2003 to 2007, whereas our larger sample allows us to estimate the degree of selection in each year. This matters because of

20 Starting in 2011 PFFS plans were required to have networks; at that time many transformed themselves into PPO plans.

21 This was by law in the case of the PFFS plans.

22 The enrollment data are from annual reports of the Medicare Payment Advisory Commission.

23 A plan could enter Part C as a "cost-based" plan in the very early years of the program, but this has not been possible for many years.

24 Risk adjustment for this group during their first year of enrollment is based just on demographic variables.

the phase in of the new risk-adjustment system from 2004 to 2007 and the changes in the lock-in policy in 2006 and again in 2007. We emphasize a comparison of years before both of these policy changes with the years after the new risk-adjustment system was fully in place, since that is the cleanest test of the new incentives.

Second, compared with Brown et al.'s (2014) sample, we add four more years at the end of the period when the new risk-adjustment system was fully implemented. Because of the policy changes within the 2003–07 period just noted, pooling those years is problematic. Moreover, pooling the years from 1994 to 2001 is also problematic because MA reimbursement policy changed markedly during that period.[25] For that reason, but also for reasons of cost and convenience, we begin our sample in 2001. Confirming the earlier studies of selection, we find substantial selection in 2001 and 2002.

Third, although Brown et al. (2014) include all MA enrollees, we limit our sample to the subgroup of MA beneficiaries who are elderly, who are not institutionalized, and who are not eligible for Medicaid ("non-duals"). Starting in 2004 Medicare introduced MA plans specifically targeted at the institutionalized and those eligible for Medicaid, many of whom are nonelderly, making it problematic to compare the earlier and later periods for those groups. The sample we study is around two-thirds of all Medicare beneficiaries in the years we include and a somewhat higher proportion of all MA enrollees.

Fourth, we estimate selection effects both including and excluding MA beneficiaries in Private Fee-for-Service (PFFS) plans. Because PFFS plans did not have provider networks before 2011 and so in fact resembled TM with enhanced benefits, any selection effects in PFFS could differ from the more common HMO and PPO type MA plans. Moreover, enrollment in PFFS plans grew rapidly starting in 2004 (McGuire, Newhouse, and Sinaiko 2011). Including PFFS plans in analyses is important, however, to the extent that TM beneficiaries switching into PFFS plans after 2004 might have switched into HMO or PPO plans before 2004.

V. Results on Plan Selection

Descriptive data on the group we analyze are shown in Table 2. Similar data on the two groups we exclude, the duals and the institutionalized, are shown in Appendix Tables 1 and 2 (see Appendix Tables 1–4 online at www.mitpressjournals/doi/suppl/10.1162/ajhe_a_0001). As is readily apparent in Appendix Tables 1 and 2, the duals and the institutionalized are both fewer and spend considerably more than the non-institutionalized, non-duals. Appendix Tables 1 and 2 also show that the number of switchers among the duals and the institutionalized is always under 25 percent of the switchers and under 20 percent in the pre-HCC years. Appendix Tables 3 and 4 disaggregate the sample shown in Table 2 into those switching to HMOs and PPOs versus those switching to PFFS.

25 In particular, the Balanced Budget Act of 1997 instituted floors on MA reimbursement in previously low-paying areas starting in 1998 and also limited annual reimbursement increases in high-paying areas to 2 percent, thus causing MA enrollment to fall and a number of MA plans to exit in the 1998–2000 period, whereas enrollment had more than doubled in the 1994–97 period (McGuire, Newhouse, and Sinaiko 2011). See also Medicare Payment Advisory Commission (2001, 2002).

TABLE 2. Unadjusted expenditures and risk scores, by year and switching status, non-institutionalized, non-duals

	Stay in TM in next year			Switch to MA in next year		
	N	Mean $	Mean risk score	*N*	Mean $	Mean risk score
2001	3,731,431	5,325	0.922	24,562	3,739	0.788
		(5.8)	(0.0004)		(55)	(0.014)
2002	3,851,378	6,001	0.953	26,129	4,006	0.804
		(6.6)	(0.0004)		(60)	(0.013)
2003	3,971,622	6,358	0.976	41,798	4,183	0.809
		(6.7)	(0.0004)		(46)	(0.018)
2004	3,927,310	6,651	0.983	77,193	4,566	0.842
		(7.0)	(0.0004)		(36)	(0.023)
2005	3,789,618	6,871	1.000	160,230	4,722	0.853
		(7.3)	(0.0004)		(26)	(0.033)
2006	3,687,693	6,958	1.018	161,753	5,750	0.927
		(7.5)	(0.0005)		(30)	(0.031)
2007	3,610,107	7,016	1.023	156,727	5,822	0.947
		(7.6)	(0.0005)		(31)	(0.031)
2008	3,531,134	7,015	1.004	126,855	5,374	0.883
		(7.8)	(0.0004)		(33)	(0.027)
2009	3,537,119	7,263	1.010	111,616	5,776	0.898
		(8.2)	(0.0004)		(39)	(0.023)
2010	3,577,950	7,343	0.967	91,620	5,590	0.846
		(8.2)	(0.0004)		(41)	(0.021)

Notes: 2007 dollars. Standard errors are in parentheses. The following types of beneficiaries were excluded from the sample: beneficiaries whose original eligibility was attributable to disability, ESRD, or Alzheimer's; newly eligible beneficiaries (since no prior claims information was available); beneficiaries who did not have 12 months of continuous enrollment in TM, both Parts A and B, in the prior year; the institutionalized; and dual eligibles. In addition, beneficiaries who switched into cost MA plans or Special Needs Plans were excluded. Risk scores are computed using the 2007 CMS-HCC model.

Two conclusions are immediately apparent from the data in Tables 1 and 2. First, as already noted, the proportion of beneficiaries who switch in any given year is small, 1 to 4 percent. Second, the raw means are suggestive of considerable selection; in the 10 years shown in Table 2, mean spending in TM among those who switched to MA in the year before they switched was 17 percent to 34 percent less than among the much larger group that remained in TM.

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The left-hand panel of Table 3 shows the results from reestimating Brown et al.'s (2014) second equation using the sample described in Table 2. The key coefficients are those for *Switched to MA the following year* and *Switched to MA the following year × year.* For those who were in TM in 2001 and switched to MA on January 1, 2002, the *Switched to MA* coefficient indicates that they spent $174 less on average within each HCC than those who remained in TM assuming they spent the full year in MA, and in 2002 $393 less (174 + 219). Using the data from Table 2, these amounts were 3–6 percent of mean spending in those years. In 2003–05, as the transition to HCCs was occurring, favorable selection within HCC increased, especially among those choosing to switch to MA in 2004 and 2005. The results for 2004 and 2005 are consistent with Brown et al.'s (2014) result that within-HCC selection increased among those choosing to enroll in MA in 2003 and later.

After 2005, however, selection fell, contrary to Brown et al.'s (2014) result. In fact, during 2006–10, the years when the HCC system was fully in place, the amount of selection within HCC was approximately the same as in 2001–02, before the HCC system was implemented. Specifically, if one averages the within-HCC selection for 2001–02, the two years before the transition began, the within-HCC selection is $284 ((174 + 393)/2) with a standard error of $99; the analogous figure for 2006–10, the five years after the HCC system was fully in place, is $320 with a standard error of $42. The $36 difference between these two means is well within one standard error of the difference.

Although the data in Table 3 suggest that in 2003–05 there was additional favorable selection within HCC when compared with the pre-HCC years, in those years there was 70 percent, 50 percent, and 25 percent weight, respectively, on the old demographic system. As a result, it may well not have been profitable for a plan to select within HCC, especially in 2003 and 2004, because during these phase-in years even low-cost patients in a high-cost HCC likely would be more expensive than a patient with no HCCs and so would have reduced the plan's profit. We therefore do not regard the data for these transition years as necessarily indicating that plans chose to select within HCC.

The right-hand panel in Table 3 shows results for the same years using the old demographic risk-adjustment methods, that is, excluding the risk score but adjusting for age and gender.[26] The results show much greater overpayment due to selection under the old system. When only age and gender are adjusted for, those who remained in TM in 2001 and 2002 spent $1,689 and $2,239 (1,689 + 550) more in TM than those who switched, respectively. Thus, the comparison of the $1,689 and $2,239 values (average $1,964) with the $320 figure from the five years following full implementation of the HCCs (Table 3, left-hand panel) indicates that the combination of the HCC risk-adjustment system and the lock-in reduced overpayment attributable to selection by roughly a factor of five.

Although Table 3 contains our principal findings, we examined the robustness of our findings to including as explanatory variables MA penetration in the beneficiary's county of residence and a main effect for those who moved zip code of residence (Table 4). The

26 The earlier risk-adjustment formula also adjusted for Medicaid and institutionalized status, but those groups are not in our sample.

TABLE 3. Outcome = expenditures in year t, predictor = fraction of year t+1 in MA (risk score adjustment vs. age-sex adjustment), non-institutionalized, non-duals

		Model 1: Adjusted for risk score		Model 2: Not adjusted for risk score	
		Coeff.	Std. err.	Coeff.	Std. err.
Fraction of months in MA in next year		−174.13	63.09	1,688.84	76.19
Year (reference = 2001)	2002	333.81	6.06	676.18	7.63
	2003	441.20	6.32	1,014.92	8.12
	2004	660.89	6.54	1,302.56	8.47
	2005	696.30	6.80	1,510.95	8.82
	2006	586.26	7.18	1,591.44	9.41
	2007	593.84	7.26	1,643.04	9.53
	2008	798.21	7.52	1,652.43	9.65
	2009	984.16	7.75	1,922.47	9.94
	2010	1,523.65	7.85	2,012.32	9.97
Fraction of months in MA in next year × year	2002	−218.87	98.75	−549.81	122.32
	2003	−289.79	86.33	−828.00	105.29
	2004	−553.46	77.99	−772.56	95.99
	2005	−513.90	68.19	−571.23	82.88
	2006	−34.95	68.20	556.52	83.47
	2007	−227.04	68.32	589.73	83.57
	2008	−147.75	69.72	242.79	84.81
	2009	−89.04	71.62	501.16	87.40
	2010	−231.84	71.95	231.72	87.81
Risk score (centered on 1)		10,829.92	7.69		
Age group (reference = 65–69)	70–74			1,039.19	8.90
	75–79			2,198.67	10.18
	80–84			2,973.63	11.05
	85–89			3,449.23	13.09
	90–94			3,628.99	18.36
	95+			3,113.58	32.25
Male				48.40	10.49
Male × age group	70–74			325.25	13.90
	75–79			819.12	16.26
	80–84			1,123.17	18.35
	85–89			1,225.91	22.44
	90–94			1,143.91	33.94
	95+			1,427.16	68.08
Constant		6,174.42	4.69	3,215.42	8.52

Notes: The sample is the same as in Table 2. Coeff. = coefficient, Std. err. = standard error.

TABLE 4. Model results with MA penetration: Outcome = expenditures in year t, predictor = fraction of year t + 1 in MA (risk score adjustment vs. no adjustment), non-institutionalized, non-duals

		Model 1: Adjusted for risk score		Model 2: Not adjusted for risk score	
		Coeff.	Std. err.	Coeff.	Std. err.
Fraction of months in MA in next year		−210.18	63.32	−1,964.87	76.49
Year (reference = 2001)	2002	638.70	13.66	1,622.14	17.90
	2003	753.38	13.65	1,966.81	17.99
	2004	1,007.14	13.83	2,293.03	18.28
	2005	1,047.99	14.29	2,529.94	18.88
	2006	964.23	15.21	2,680.92	20.14
	2007	1,018.21	15.98	2,745.02	21.16
	2008	1,287.94	16.82	2,765.90	22.00
	2009	1,502.92	17.61	3,038.70	22.93
	2010	2,165.28	17.69	3,163.19	22.90
Fraction of months in MA in next year × year	2002	−266.21	99.04	−617.10	122.70
	2003	−308.93	86.61	−811.81	105.68
	2004	−526.14	78.23	−641.32	96.30
	2005	−480.42	68.41	−345.08	83.17
	2006	−0.69	68.41	826.28	83.74
	2007	−181.04	68.54	850.10	83.87
	2008	−91.17	69.96	501.91	85.11
	2009	−25.10	71.85	757.69	87.70
	2010	−160.31	72.16	491.97	88.10
MA penetration in county (10% increase)		23.92	3.38	182.67	4.27
MA penetration in county (10% increase) × year	2002	25.96	4.71	26.25	5.93
	2003	13.71	4.99	9.70	6.42
	2004	−21.66	5.16	−35.60	6.70
	2005	−24.84	5.33	−73.33	6.90
	2006	−41.65	5.47	−142.98	7.14
	2007	−64.49	5.74	−152.11	7.46
	2008	−91.37	5.83	−157.29	7.43
	2009	−101.18	6.04	−163.35	7.65
	2010	−156.00	5.96	−179.84	7.52

TABLE 4. *Continued*

		Model 1: Adjusted for risk score		Model 2: Not adjusted for risk score	
		Coeff.	Std. err.	Coeff.	Std. err.
Moved		334.30	11.51	961.84	15.32
Risk score (centered on 1)		10,829.43	7.70		
Age group (reference = 65–69)	70–74			1,043.19	8.90
	75–79			2,196.12	10.18
	80–84			2,961.12	11.05
	85–89			3,428.98	13.09
	90–94			3,604.20	18.35
	95 +			3,091.45	32.22
Male				44.28	10.50
Male × age group	70–74			328.06	13.90
	75–79			826.37	16.25
	80–84			1,132.36	18.34
	85–89			1,233.97	22.42
	90–94			1,148.87	33.91
	95 +			1,428.28	68.03
Constant		5,810.87	12.97	2,034.76	18.15

Notes: The sample is the same as Table 2. Coeff. = coefficient, Std. err. = standard error.

inclusion of these variables did not change our qualitative conclusions, namely that at a given level of penetration, selection in the era before HCCs with only demographic adjusters was notably greater than in the era after implementation of the HCCs, and that selection within HCCs was little different among those who switched in the 2006–10 period than it was in 2001 and 2002.

Interestingly, however, after 2002 the effect of a 10 percentage point change in MA penetration in a county became monotonically more negative. This was a period when MA enrollment was rising substantially, from around 15 percent of all beneficiaries in 2001–02 to around 24 percent in 2010–11, presumptively because the statutory MA reimbursement formula was raising MA reimbursement relative to TM (McGuire, Newhouse, and Sinaiko 2011). The usual economic models of selection described above suggest that an exogenous increase in MA penetration should attract worse risks into MA and thus increase the average MA risk score. The data in Table 4 do not support this prediction; starting in 2005 the coefficient on the penetration variable turns negative and becomes steadily more negative irrespective of whether risk score is included in the risk-adjustment method. During this period MA penetration rose each year; the steadily more negative coefficients indicate that, if anything, increasingly better risks were being attracted to MA as county-level penetration increased in the cross section.[27]

In addition to offering plans an incentive to select within HCC, the HCC risk-adjustment scheme offers plans an incentive to code diagnoses as completely as possible, since additional recorded diagnoses can increase a plan's risk-adjusted payment. Unlike the incentive to select within HCC, plans appeared to act on the incentive to code more intensively. To account for more intensive coding, the Centers for Medicare and Medicaid Services reduced risk scores and hence reimbursement for MA plans 3.4 percent in 2010, 2011, and 2012 (Medicare Payment Advisory Commission 2012). Any more intensive coding by plans over time, however, should not have a direct effect on our estimates, since all of our data come from beneficiaries in TM. In theory, however, there could be a spillover effect from more intensive coding in MA such that counties with greater proportions of Medicare beneficiaries in MA might also code more intensively for TM beneficiaries. The results in Table 4, however, do not support this hypothesis since controlling for MA penetration does not change our conclusions.

As described above, characteristics of PFFS plans differed considerably from those of HMO and PPO plans, and one might therefore anticipate selection patterns would differ. Hence, we reestimated the results in Table 3 using only those who switched to HMO or PPO plans.[28] These results are shown in Table 5. We were unable to identify whether the plan was a PFFS or an HMO or PPO plan for those who switched in 2010, so Table 5 has one less post-HCC year than Table 3. The average switcher in the four post-HCC years shown in Table 5 spent $372 less in TM the year prior to the switch, compared with the $320 figure in Table 3. This difference of $52 is also well within one standard error; thus, our findings are robust to considering just those who switch to HMO and PPO plans rather than all MA plans, including PFFS.

We suggested above that the more unique physicians a beneficiary anticipated seeing, the more likely the beneficiary would prefer TM because it would be more likely that they would want to use out-of-network physicians if they enrolled in MA. Table 6 shows that the average number of unique physicians seen by a beneficiary rises monotonically with the number of HCCs. Similarly, the costliest beneficiaries within an HCC may prefer the inherently greater provider choice in TM.

VI. Conclusions

We have two main results. First, although there appears to be selection within HCC that favors MA, the magnitude of such selection changed little from the period two years before the HCC system was introduced to the five years following its full introduction. Second, though some favorable selection into MA appears to remain, the introduction of the HCC

27 In Newhouse et al. (2012) we found that risk scores among those switching to MA had no consistent relationship with penetration levels in the cross section over the 2004–08 period, which is also not consistent with the usual economic models of selection.

28 Very few persons switched to PFFS in 2001 and 2002 (Online Appendix Table 4, www.mitpressjournals/doi/suppl/10.1162/ajhe_a_0001) so that comparisons of pre- and post-HCC years for just PFFS plans are imprecise.

TABLE 5. Outcome = expenditures in year t, predictor = fraction of year t + 1 in MA (risk score adjustment vs. no adjustment), non-institutionalized, non-duals switching to HMOs and PPOs

		Model 1: Adjusted for risk score		Model 2: Not adjusted for risk score	
		Coeff.	Std. err.	Coeff.	Std. err.
Fraction of months in MA in next year		−196.69	63.32	−1,704.53	76.56
Year (reference = 2001)	2002	337.20	6.06	676.46	7.63
	2003	446.83	6.32	1,016.29	8.12
	2004	667.62	6.54	1,305.89	8.47
	2005	704.61	6.80	1,514.82	8.82
	2006	597.28	7.18	1,594.98	9.42
	2007	604.86	7.26	1,645.76	9.53
	2008	807.31	7.52	1,654.99	9.66
	2009	993.77	7.77	1,924.48	9.94
Fraction of months in MA in next year × year	2002	−211.06	99.51	−507.30	123.54
	2003	−355.56	88.45	−808.93	108.16
	2004	−659.68	81.88	−585.88	101.25
	2005	−683.76	72.11	−486.88	87.82
	2006	−196.44	76.14	−55.24	94.60
	2007	−155.84	73.67	194.29	90.71
	2008	−228.47	73.21	−160.82	89.11
	2009	−120.70	74.07	390.32	90.71
Risk score (centered on 1)		10,724.59	7.93		
Age group (reference = 65–69)	70–74			1,023.49	9.34
	75–79			2,161.12	10.55
	80–84			2,904.99	11.43
	85–89			3,344.25	13.52
	90–94			3,485.41	18.98
	95+			2,966.50	33.10
Male				67.11	11.01
Male × age group	70–74			327.83	14.62
	75–79			830.72	16.90
	80–84			1,132.71	18.97
	85–89			1,235.56	23.27
	90–94			1,142.96	35.21

TABLE 5. *Continued*

	Model 1: Adjusted for risk score		Model 2: Not adjusted for risk score	
	Coeff.	Std. err.	Coeff.	Std. err.
Male × age group, *continued*				
95+			1,357.56	70.25
Constant	6,166.15	4.70	3,246.75	8.72

Notes: 2007 dollars. The following types of beneficiaries were excluded from the sample: beneficiaries whose original eligibility was attributable to disability, ESRD, or Alzheimer's; newly eligible beneficiaries (since no prior claims information was available); beneficiaries who did not have 12 months of continuous enrollment in TM, both Parts A and B, in the prior year; the institutionalized; and dual eligibles. In addition, beneficiaries who switched into PFFS, cost MA plans or Special Needs Plans were excluded. MA-PPO enrollees are included. Values for 2010 are not shown in this table because we are unable to identify the type of MA plan that was chosen in 2010. Coeff. = coefficient, Std. err. = standard error.

TABLE 6. Mean number of providers seen in 2011 by number of HCCs

No. of HCCs	Subjects	Mean no. of providers	SD	Min	Max
0	1,642,618	3.50	3.74	0	90
1	1,091,932	5.38	4.48	0	101
2	633,884	6.74	5.33	0	104
3	346,879	8.00	6.18	0	101
4	406,325	10.55	8.31	0	225

Source: 20% sample of 2011 Medicare claims sample and the Carrier file. Providers include MDs, DOs, and RNs billing independently other than radiologists, anesthesiologists, and pathologists as identified from the provider specialty variable on the Carrier file. All differences are significant at the 0.0001 level.

system and the lock-in has markedly diminished it. Since these two innovations were introduced at approximately the same time, we have not tried to decompose the amount of reduction attributable to each. This decrease in selection is consistent with other evidence (McWilliams, Hsu, and Newhouse 2012; Newhouse et al. 2012; Newhouse et al. 2013; Newhouse and McGuire 2014).

A key policy question, of course, is the magnitude of any remaining overpayment due to selection in the MA program. How much should our findings influence one's judgment on that question? Given the small proportion of beneficiaries who switch in a given year relative to the stock of enrollees in TM, 1–4 percent of all beneficiaries, we think rather little. Beneficiaries tend to remain in the program they selected, and over time there is regression to the mean. Stated differently, the longer that beneficiaries remain in TM or MA, the more that concerns about the effects of any initial differential selection are mitigated, as the mortality data cited above indicated. Moreover, the comparison of self-rated health status among all TM and MA beneficiaries also cited above found an insignificant difference by 2007. Thus, the mix of risks among all those enrolled in MA and TM (excluding the duals and the institutionalized) may be close.

We return now to the two types of selection identified in the health economics literature and the associated inefficiencies, and consider what our results mean in those terms. The first type of selection occurs when plans distort their mix of premiums and benefits to attract profitable enrollees. The new risk-adjustment system changed plans' financial incentives to favor low-cost beneficiaries within an HCC rather than low-cost beneficiaries within an age-sex group. Our first conclusion suggests that plans did not alter their selection behavior in response to this change in incentive. We find this conclusion plausible; while plans choose networks, formularies, cost-sharing structure, and marketing, under the new system they would have to motivate the physicians and hospitals with whom they contract to select profitable patients *within an HCC*. It is not obvious how plans, which mostly have arms-length, fee-for-service contracts with hospitals and physicians, could do that. It certainly is conceivable that a plan could form networks of providers who happen to have low-cost patients within a diagnosis at a given point in time, but given the randomness in spending among a given panel of patients at the provider level along with the turnover in patients at any given provider or provider group, providers that appear to be low-cost today may be high-cost tomorrow (Hofer et al. 1999; Bronskill et al. 2002). Also the more stringent lock-in periods may have inhibited some plan switching in response to health status changes.

The conclusion that plans did not alter their selection mechanisms to select within HCCs is consistent with the finding from other work that failed to find evidence of selection across HCCs (Newhouse et al. 2013). It seems plausible that it is less costly for plans to select across HCCs than within HCCs, since networks and formularies might be structured to attract or not attract individuals with certain diagnoses. Our failure to find this type of selection with the HCC system in place strengthens our confidence that selection mechanisms did not change in response to the HCCs in such a way as to attract low-cost beneficiaries within an HCC.

The second type of selection problem occurs when individuals sort themselves inefficiently between fixed plan types, due to the premium for the more generous plan being "too high" because more costly types are more likely to prefer the more generous coverage. The premium thus reflects both the average incremental cost of the plans' differing generosity (as it needs to for efficiency) and also a component due to selection (which interferes with efficient sorting). When risk adjusters better track expected costs, premium differences between plans in a competitive market will contain less of a component due to selection.[29] Thus, our findings suggest that inefficiencies in plan choices in MA have

29 Although premiums and cost sharing for TM are set administratively, the expected cost of MA premiums and cost sharing is constrained to be equal to or less than TM's actuarial value. In practice by 2007 the expected cost of MA was less for almost all beneficiaries (McWilliams et al. 2011). Thus, this constraint on the actuarial value of MA plans is rarely binding.

likely diminished. It is worth keeping in mind, however, that any single premium cannot sort beneficiaries efficiently, so some selection problems would remain even if selection as we measure it here were completely eliminated (McGuire et al. 2013). Efficient plan choice of premium and benefits may also be interfered with by institutional rules that discourage plans from deviating from a premium exactly equal to the mandatory Part B premium automatically paid by all beneficiaries who elect an MA plan (Newhouse and McGuire 2014).

One other phenomenon may explain some remaining selection. Newhouse et al. (2012) showed that more than half of those who disenrolled from Medicare Advantage reenrolled within 12 months from disenrolling. This is consistent with beneficiary provider shopping; when beneficiaries want a medical procedure performed by an out-of-network provider, for example, to have a hip replacement done by a particular surgeon, they may switch to TM to have it done, and then reenroll in the MA plan. Although the lock-in should have reduced such provider shopping, this type of behavior may plausibly have occurred both before and after the HCC system and the lock-in, that is, it is related to the inherent difference between TM and MA that TM freedom-of-provider choice.

Balancing profitability across groups of beneficiaries according to their health-care use, as the HCC system seeks to do, is likely to mitigate incentives to distort certain services in relation to others, but our findings do not allay concerns about the basic incentives to reduce premiums and benefits and structure networks and formularies to attract profitable types. Nonetheless, in contrast to the basic Rothschild-Stiglitz model where there is no regulator and no constraint on contracts that can be offered, Medicare has numerous instruments with which to address selection, and it appears to have made substantial progress in doing so.

In sum, improving the match between risk-adjusted payments and expected costs in MA is likely to have improved the efficiency of the MA program, both in terms of efficient sorting of beneficiaries between TM and MA and in terms of the nature of plan offerings. Efficiency problems would, however, persist even with a complete elimination of selection as studied here because of Medicare's single premium policy. Attention to premium policy as well as risk-adjustment policy is called for to more fully address performance of the MA program.

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